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Money, Output, and Prices: Evidence From a New Monetary Aggregate

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This article derives a new utility-based monetary aggregate, the currency-equivalent (CE) aggregate. It equals the stock of currency that would be required for households to obtain the liquidity services that they get from their entire collection of monetary assets. This aggregate is derived from preferences assuming that these satisfy a separability assumption in addition to satisfying the requirements for Divisia aggregation. The resulting aggregate remains valid when asset characteristics change and equals the sum of individuals’ CE holdings. It also predicts output movements better than simple-sum aggregates such as M1 and M2.

KEY WORDS: Aggregation; Divisia aggregation; Monetary policy; Money-income causality.

Conventional simple-sum monetary aggregates, such as the monetary base, M1, and M2, arbitrarily assign equal weight to all of the component assets included in the aggregate and zero weight to all other assets. A more attractive approach to monetary aggregation involves weighing different assets by the value of the monetary services they provide. This principle underlies Barnett’s (1980) derivation of Divisia monetary aggregates. The continued widespread use of conventional simple-sum aggregates in spite of their weak theoretical foundation is particularly surprising because research has repeatedly shown that Divisia aggregates are at least as good as predicting gross national product (GNP) (Barnett, Fisher, and Serletis 1992; Serletis 1988). In this article, we derive and study the properties of a related aggregate, which we call the currency-equivalent (CE) monetary aggregate. This aggregate, which was proposed but not analyzed in depth by Hutt (1963) and Rotemberg (1991), is a time-varying weighted average of the stocks of different monetary assets, with weights that depend on each asset’s yield relative to that on a benchmark “zero liquidity” asset.

The CE aggregate can be interpreted as the stock of currency that yields the same transactions services as the entire constellation of monetary assets. Currency has a weight of unity in the CE aggregate, whereas other assets with higher rates of return receive a smaller weight because their marginal liquidity services are lower. In this article, we first show that the CE aggregate can be given a preference-theoretic foundation under conditions that are slightly more stringent than those that are needed to give such foundations to Divisia aggregates. We then use this aggregate to develop new evidence on the correlation between money and output.

This article is divided into seven sections. Section 1 derives the CE monetary aggregate and compares it with existing Divisia aggregates. Section 2 describes the data inputs we use to construct the CE aggregate for the United States and reports summary statistics on its univariate time series properties. Section 3 describes our methodology for studying how monetary shocks affect output and prices. Section 4 presents tests of the CE aggregate’s ability to Granger-cause real output and compares its predictive performance to that of other indicators of monetary policy. It also tests the long-run neutrality of money with respect to output. Section 5 presents a parallel analysis of the effect of monetary shocks on prices, comparing the findings for the CE aggregate with those for more traditional measures of monetary policy. Section 6 evaluates the robustness of our bivariate results on money, output, and prices by considering a vector autoregressive (VAR) model. As in the paper by Sims (1980), the VAR also includes the nominal interest rate. Section 7 concludes and sketches the implications of our findings for the estimation of money demand and for competing theories of the monetary transmission mechanism.

1. THE CURRENCY-EQUIVALENT MONETARY AGGREGATE: DERIVATION AND PROPERTIES

Some assets can readily be used for transactions. Individuals pay for the liquidity that these assets offer by forgoing the higher expected returns that are available on other, less liquid assets. Holding one dollar in currency costs more than holding one dollar in a negotiable order of withdrawal (NOW) account, and it presumably generates greater liquidity, or transactions, services. We formalize this
idea by assuming that individuals derive different amounts of utility from holding different assets. Our results could also be obtained by assuming that individuals and firms incur transactions costs that depend on their portfolios of assets and on the volume of transactions, as shown by Feenstra (1986).

1. Deriving the CE Aggregate

We consider an individual whose expected lifetime utility in period $t$ is given by

$$V = E_t \sum_{j=0}^{\infty} \beta^j U(C_{nj}, L_{nj}),$$

where $E_t$ takes expectations as of time $t$, $\beta$ is an intertemporal discount factor, $C_t$ is consumption, $L_t$ is the aggregate of liquidity services, and $U$ gives the level of instantaneous utility. We assume that this function $U$ is concave in both arguments. The aggregate of liquidity services is given by

$$L_t = f(m_{0,t}, m_{1,t}, \ldots, m_{n-1,t}, \alpha_t),$$

where $m_{i,t}$ represents the amount of asset $i$ held at $t$ and $m_{0,t}$ denotes the amount of currency held at $t$. Unfortunately assets do not have unchanging physical characteristics, so the level of liquidity services they provide varies over time. This is captured by the time-varying parameter $\alpha_t$. This parameter would change and thereby change the function $f$ if, for example, there is a change in the number of checks that can be written on a particular account.

For many purposes, it is desirable to construct monetary aggregates that equal (or at least approximate) $L$. This is of particular interest if one feels that it is the amount of money that enters the utility function that is related to aggregate demand or if one wants to estimate a money-demand relationship. We now consider assumptions on the aggregator function $f$ that allow us to recover $L_t$.

First, we assume that for every value of $\alpha$ this aggregator is homogeneous of degree 1 in all its monetary arguments. With this assumption, the Divisia aggregate can be used to measure the percentage change in $L$. If $L_t$ were not homogeneous of degree 1 in its arguments, we could follow Caves, Christensen, and Diewert (1982) and focus on the distance function $d(m_{0,t}, m_{1,t}, \ldots, \alpha_t, L_0)$, which gives the amount by which all monetary assets must be divided to obtain a subutility of monetary assets equal to $L_0$. By construction, this distance function is homogeneous of degree 1 in monetary assets.

Second, we assume that $f$ is additively separable in currency ($m_{0,t}$) and other monetary assets:

$$f(m_{0,t}, m_{1,t}, m_{2,t}, \ldots) = h(m_{0,t}) + k(m_{1,t}, m_{2,t}, \ldots, \alpha_t).$$

This assumption gives a central role to currency because it is possible to obtain any level of liquidity services by holding sufficient currency. Although the inessentiality of other monetary assets might be controversial, it is consistent with the fact that circulating media of exchange appear to predate the introduction of other liquid assets.

When coupled with the homogeneity of $f$, (3) represents a strong assumption. In particular, it then implies that $L$ can be normalized so that

$$L_t = f(m_{0,t}, m_{1,t}, m_{2,t}, \ldots, \alpha_t) = m_{0,t} + k(m_{1,t}, m_{2,t}, \ldots, \alpha_t).$$

To see this, note first that the linear homogeneity of $f$ implies that

$$L_t = \sum_{i=0}^{n-1} f_{i,t} m_{i,t},$$

where $f_{i,t}$ is the partial derivative of $f$ with respect to the quantity of monetary asset $i$ evaluated in period $t$. Since (3) implies that $f_{0,t} = h_{0,t}$,

$$h(m_{0,t}) = h_{0,t} m_{0,t} = f_{0,t} m_{0,t}.$$
the utility function is separable in consumption, \( L \), and these other services.

The benchmark return has the standard property that

\[
U_c(C_t, L_t) = (1 + r_{b,t})E_t \frac{P_t \beta U_c(C_{t+1}, L_{t+1})}{P_{t+1}},
\]

where \( P_t \) measures the price of a unit of consumption in terms of currency at date \( t \). Equation (6) says that consumers are indifferent between their current (optimal) consumption profiles and profiles in which they increase their holdings of the benchmark asset by reducing current consumption and use the proceeds to increase future consumption.

Monetary assets do not share this property of the benchmark asset because increasing their stock raises the level of liquidity services received by the consumer. Thus the loss in utility from lowering current consumption is offset by both mark asset because increasing their stock raises the level of liquidity services received by the consumer. Thus the loss in utility from their reduced holdings of monetary asset \( i \) is held in strictly positive amounts, it must be the case that

\[
U_c(C_t, L_t) = U_c(C_t, L_t)f_{i,t} + (1 + r_{i,t})E_t \frac{P_t \beta U_c(C_{t+1}, L_{t+1})}{P_{t+1}}.
\]

Substituting the expectational term from (6) into (7), we obtain

\[
U_c f_{i,t} = \frac{r_{b,t} - r_{i,t}}{1 + r_{b,t}} U_c.
\]

This says that the consumers are indifferent between reducing their holdings of monetary asset \( i \) by one unit and increasing their holdings of the benchmark asset by \((1 + r_{i,t})/(1 + r_{b,t})\). This change leaves the value of the individual’s portfolio at the beginning of period \( t + 1 \) unaffected so that future consumption can remain unaffected as well. This portfolio change, however, allows individuals to raise their current consumption by \((r_{b,t} - r_{i,t})/(1 + r_{b,t})\). The gain in utility from this consumption increase must therefore equal the fall in utility from their reduced holdings of monetary asset \( i \).

The issue at this point is whether consumers will actually choose to hold strictly positive amounts of currency and of the other monetary assets or whether they will set either currency or the aggregator \( k \) equal to 0. This issue arises because (4) is linear in these two quantities. This implies that, unless the relative prices ensure that the cost per unit of liquidity provided by these two quantities is equal, the consumer will set one of these to 0. We thus analyze now the condition on relative prices that leads consumers to an interior solution.

An interior solution obtains if the cost in terms of foregone consumption of obtaining one unit of \( k \) is the same as the cost of obtaining one unit of \( m_0 \). The cost of obtaining one unit of \( k \) is the result of minimizing

\[
\sum_{i=1}^{n-1} \frac{r_{b,t} - r_{i,t}}{1 + r_{b,t}} m_i,
\]

with respect to \( m_i \) subject to the constraint that \( k \) is equal to 1. An interior solution obtains only if the minimized value of (9) is equal to the cost of holding one unit of currency; namely, \( r_{b,t}/(1 + r_{b,t}) \).

This shows that the assumption that the solution is interior is not innocuous. If tastes change in a way that changes the \( f_{i,t} \)'s (as would occur if the liquidity services of all monetary assets other than currency went up), interest-rate differentials would have to change for this equality to be satisfied. If the interest differentials did not change, such a change in tastes would lead individuals to choose a corner solution in which either currency or the other monetary assets would be absent from their portfolios. The linear structure of the aggregator in (4) implies that our interiority assumption is satisfied only if interest rates do respond to changes in tastes.

Using the normalization \( f_{i,t} = 1 \) and maintaining the assumption of an interior solution, the analog of (8) for currency is

\[
U_c = \frac{r_{b,t}}{1 + r_{b,t}} U_c.
\]

Combining Equations (8) and (10) now yields

\[
f_{i,t} = \frac{r_{b,t} - r_{i,t}}{r_{b,t}}.
\]

The level of the liquidity aggregate \( L_t \) from Equation (4) then satisfies

\[
L_t = \sum_i \frac{r_{b,t} - r_{i,t}}{r_{b,t}} m_i \equiv CE_t.
\]

This expression defines the CE monetary aggregate. The weight of each asset in this aggregate is \((r_{b,t} - r_{i,t})/r_{b,t}\). In light of Equation (11), this weight equals the marginal utility of asset \( i \) relative to that of currency.

Our derivation has considered the level of liquidity held by a person. Because (12) is linear in individual asset holdings, however, the sum of the \( L \)'s held by all individuals is simply (12) applied to aggregate asset holdings. Thus \( CE_t \) provides an accurate measure of the sum of the individual \( L \)'s even if the aggregator functions \( f \) differ for different individuals.

The CE aggregate has the attractive property that assets that do not pay interest, such as currency and travelers' checks, are added together with weights of unity. Other assets are added with weights between 0 and 1, with higher-yield assets receiving lower weights. This makes intuitive sense because these assets with higher returns must provide smaller liquidity services. Assets whose expected return exceeds that on the benchmark asset ought to be given a 0 rather than a negative weight. Equation (6) should be satisfied for these assets, so their higher expected return should simply be a compensation for risk.

The CE aggregate also adapts easily to changes in the financial environment. When new assets, such as interest-bearing NOW accounts, are introduced, they can be added to the aggregate. The problem of deciding whether such accounts are "liquid enough" to be included can be resolved by reference to their interest yield.

One interesting interpretation of this aggregate was developed by Barnett (1991). He showed that, assuming static expectations, \( CE_t \) equals the expected present discounted value of expenditures on liquidity services, where those expenditures can be measured using the Divisia index. The expen-
ditutes in period $t$ equal $\sum_i (r_{i,t} - r_{i,0}) m_{i,t}$, which is the total amount of interest foregone by holding monetary assets instead of the benchmark asset. Under static expectations, the present value of these expenditures equals their level divided by the rate of return on the benchmark asset. This also yields the CE aggregate.

### 1.2 Properties of Divisia and CE Aggregates

Divisia aggregation requires weaker assumptions on the aggregator function $f$ for a constant $\alpha$. Bernett and Yue (1992) showed that in continuous time, assuming only that $f$ is homogeneous of degree 1 and $\alpha$ is constant, the rate of change of the Divisia index equals the rate of change of $L_t$. Deriving the CE aggregate requires an additional assumption on $f$—namely, that the services from currency be separable from those provided by other assets.

In part because of this stronger assumption, the CE aggregate can account for changes in $\alpha$, but the Divisia aggregate cannot. To see the effect of changes in $\alpha$, we totally differentiate (2) and obtain

$$dL_t \approx \text{sum over } a \left( f \left( \frac{d \alpha}{\alpha} \right) + f \right)$$

Using (4) and (10), which are valid under linear homogeneity even when separability is violated, (13) implies that

$$\frac{dL_t}{L_t} = \sum_{i=0}^{n-1} \frac{(r_{i,t} - r_{i,0}) m_{i,t}}{\sum_{j=0}^{n-1} (r_{j,t} - r_{j,0}) m_{j,t}}\frac{dm_{i,t}}{m_{i,t}} + f \frac{d\alpha}{\alpha}$$

The first term on the right side of (14) is the Divisia index, which, obviously, equals the change in $L$ only in periods when $\alpha$ does not change. By contrast, the CE aggregate incorporates changes in $\alpha$ when they occur.

Changes in the liquidity services provided by different assets, represented by changes in $\alpha_i$, constitute one of the most important challenges of monetary aggregation because the characteristics of available monetary assets change constantly. When NOW accounts eliminated charges on checks, or when automated teller machines made funds in these accounts more accessible, their liquidity properties, and therefore their respective $f_i$, changed. Divisia aggregation assumes that, so long as asset holdings do not change, the utility provided by each asset remains immutable. The CE aggregate deals with changes in asset characteristics by incorporating the idea that for asset holdings not to change $(r_{i,t} - r_{i,0})/r_{i,t}$ must rise as much as $f_i$. The CE aggregate thus interprets increases in $(r_{i,t} - r_{i,0})/r_{i,t}$ with an unchanged asset stock as an increase in the assets’ liquidity services. One consequence of this is that changes in interest rates imply that CE changes even when asset holdings do not change.

We are not suggesting that Divisia indexes are totally unable to deal with changes in the aggregator function $f$. Caves et al. (1982) showed that, even with an aggregator that changes over time, the Malmquist approximation to the Divisia has some desirable properties. It equals the product of two Malmquist indexes, the first of which gives the percent by which assets in period $t + 1$ have to be shrunk to give the same liquidity as those in period $t$ (using the aggregator for $t$) and the second of which gives the percent by which assets in period $t$ must be expanded to give the same liquidity services as those of $t + 1$ (using the aggregator for $t + 1$). The problem is that, as Equation (14) demonstrates, this product of Malmquist indexes is not very closely related to changes in $L$ itself, even when $L$ is homogeneous both before and after the change in $\alpha$.

The ability of CE to handle changes in $\alpha$ also facilitates comparisons of money holdings across individuals, states, and nations. Nations differ in their financial institutions, so their aggregators surely differ. The aggregators for different individuals are also likely to differ. Thus the CE aggregate’s ability to yield comparable levels of liquidity even when tastes and the characteristics of assets change is a benefit.

Another benefit from our separability assumption is that CE is an index that can in principle be computed with observations at a single point in time. One can meaningfully talk about the level of the CE aggregate. Even when, as is more typically the case, one is interested in the growth rate of money, this implies that the quality of its measurement does not depend on having access to observations that are close together. By contrast, Divisia aggregation requires that one use the discrete observations that are available to approximate the time derivatives of asset holdings. The quality of this approximation thus depends on the frequency with which assets are measured. Since the stock of monetary assets is sampled weekly, this approximation error of Divisia aggregates is likely to be small.

A final advantage of the CE aggregate, which we hope will provide the basis for future research, is that it simplifies the estimation of money-demand equations. Equation (10) can be viewed as a conventional money-demand equation, relating an individual’s demand for total liquidity, $L_t$, to the “market” interest rate and a measure of his or her transactions, in this case consumption. This equation applies directly to the sum of the $L_t$’s across individuals if either all individuals have identical $L_t$’s and $C$’s or if each individual’s utility function is quadratic with the same parameters. Estimates of money-demand equations often use the commercial-paper rate, which to some degree justifies our use of commercial paper as our benchmark asset. Conventional money-demand equations, however, often add other interest rates as well. Our restriction on $f$ implies the testable proposition that other interest rates have no role in the estimation of (10).

The CE aggregate also has a potentially serious disadvantage relative to Divisia indexes. The problem arises when, as seems reasonable, people do not continually rebalance their portfolios in response to interest-rate changes so that Equation (8) does not always hold. Such interest-rate changes affect the level of CE even if people keep their portfolios unchanged. By contrast, the Divisia aggregate changes only as people change their portfolios. In the case of slow responses of asset holdings to rates of return, the Divisia is thus more accurate unless one dampens the effect of these
interest-rate changes on the CE weights. We discuss this dampening later.

In spite of the differences noted previously, Divisia and CE aggregates have many things in common. Changes in monetary assets affect both aggregates only to the extent that their rate of return is lower than that of a benchmark asset. Thus changes in broker–dealer money-market mutual funds and in time deposits at commercial banks, which are in M2 but essentially yield the commercial-paper rate, do not have a significant impact on either aggregate. This contrasts with the conventional wisdom, which holds that money-market accounts can be used for making payments and so must be "like" checking accounts.

The conventional view confuses average and marginal liquidity services. The first few dollars invested in broker–dealer money-market accounts do provide liquidity services because individuals are likely to write checks on them. Because these accounts pay interest rates equal to those on zero-liquidity assets, however, they also attract additional funds that will not be used to make payments in the near future. Starting from the equilibrium holdings, changes in these asset stocks therefore do not affect total liquidity services.

One difficulty with this argument is that not all consumers start by holding the optimal portfolio of all assets. For those who do, increases in their holdings of high-yielding assets such as money-market accounts do not raise their liquidity. Other consumers are only slowly drifting into such assets. Thus they still gain inframarginal liquidity services as they acquire these assets. As these consumers reduce their "regular" checking account holdings while simultaneously raising their money-market accounts by the same amount, their liquidity does not fall as the CE and Divisia aggregates would imply. On the other hand, if the conversion from "regular" checking to money-market accounts is gradual, this has little effect on money's monthly growth rate, regardless of the form of aggregation that is used.

### 1.3 Slow Portfolio Adjustment

At least since the work of Baumol (1952) and Tobin (1956), theories of money demand have recognized that bank-transactions costs are important determinants of the average level of money held. These same transactions costs presumably slow down the adjustment of the level of money balances to changes in returns, and this led Barnett (1980) to modify the Divisia aggregate. The effect of transactions costs is that the first-order condition (8) will not hold at every point in time. Variations in the spread \( (r_{h} - r_{c})/r_{h} \) will therefore not necessarily correspond to variations in \( f \). Furthermore, given the comparatively sluggish behavior of bank deposit rates, which is partly due to regulations on the interest rates that can be paid, the error term in (8) is likely to be highly correlated with the benchmark. This complicates economic inference, as we shall discuss in more detail later.

With occasional adjustment of portfolios, individuals who are currently changing their asset stocks will choose a portfolio of monetary assets such that each \( f \) is close to the \( (r_{h} - r_{c})/r_{h} \) expected to prevail in current and future periods. The typical \( f \) in the population will therefore depend not only on current interest rates but also on past and expected future rates. Finding the exact form of this dependence would require a more elaborate theory of money demand than we have developed. For the moment we simply experiment with CE aggregates in which the weights of different assets do not just depend on current interest-rate spreads but also depend in an ad hoc way on past and future spreads. In particular, we let the weights equal centered moving averages of these spreads.

We use two approaches for constructing the future spreads that we use to compute current weights. The first assumes perfect foresight and uses actual future spreads. The two aggregates we create using this procedure are called CE-1 and CE-3, and they use 13-month and 37-month centered moving averages, respectively. At the beginning and end of the data sample, when symmetric centered moving averages cannot be computed, we use uncentered 13- and 37-month moving averages. Since the inclusion of future interest rates in constructing current monetary aggregates may bias upward the correlation of these aggregates with future variables, we also try replacing the actual future spreads with forecasts that are based on currently available information. We obtain the forecasts by fitting an autoregressive integrated moving average (ARIMA) model for the spreads over the whole sample. Our ARIMA models are second-order autoregressions as determined by the Schwarz–Bayes information criterion (SBIC).

In each month we compute a series of \( k \)-step-ahead forecasts, where \( k \) ranges from 1 to \( (n - 1)/2 \) and \( n \) is the number of months over which we smooth the spreads. This still uses future information but only to the extent that such information affects the parameter estimates in the ARIMA model. We present two such aggregates, ARIMA CE-1 (13 months) and ARIMA CE-3 (37 months).

We also considered two alternatives to our moving average procedure. The first is to use a backward moving average of the spreads. The problem with this procedure is that people changing their portfolios should look forward and not backward. Such an aggregate is thus unlikely to approximate one that solves the theoretical problems mentioned previously. A second alternative is to use a fixed-weight CE that replaces the spreads by their sample averages. Although we do present results for this aggregate, its disadvantage is that it fails to take into account changes in the liquidity characteristics of the monetary assets.

### 2. DATA AND SUMMARY STATISTICAL PROPERTIES

We include eight component assets in the CE aggregate. Three—currency (CU), travelers' checks (TC), and demand deposits (DD)—do not yield interest. The interest-paying assets are other checkable deposits (OCD), such as NOW accounts, savings accounts at thrift institutions (SSL), savings accounts at commercial banks (SCB), money-market accounts at commercial banks (MMCB), and money-market accounts at thrift institutions (MMSL). Using these symbols...
as subscripts to identify returns, the CE aggregate is given by

\[
\text{CE}_t = \text{CU}_t + \text{TC}_t + \text{DD}_t + \left(1 - \frac{r_{\text{OCD},t}}{r_{\text{CP},t}}\right)\text{OCD}_t
+ \left(1 - \frac{r_{\text{SSL},t}}{r_{\text{CP},t}}\right)\text{SSL}_t
+ \left(1 - \frac{r_{\text{SCB},t}}{r_{\text{CP},t}}\right)\text{SCB}_t
+ \left(1 - \frac{r_{\text{MMCB},t}}{r_{\text{CP},t}}\right)\text{MMCB}_t
+ \left(1 - \frac{r_{\text{MMSL},t}}{r_{\text{CP},t}}\right)\text{MMSL}_t.
\]

(15)

We include these components in the CE aggregate because they comprise the set of assets that have been traditionally considered to be monetary (i.e., those that are included in the broadest conventional definition of money, \(L\)) and that have a rate of return less than that of the benchmark rate. An alternative approach to determining the composition of monetary aggregates, pursued by Swofford and Whitney (1988), is to determine whether there is any substitutability function of monetary assets that is weakly separable from consumption and leisure. Using a nonparametric test based on the generalized axiom of revealed preference, they found that an aggregate consisting of currency, demand deposits, other checkable deposits, and savings deposits satisfies the requirements for weak separability. With the exception of money-market accounts, which were too new to be included in that study, the composition of this asset is the same as that of the CE aggregate.

Data on the quantities of various monetary assets that enter Equation (15), as well as the commercial-paper rate, are drawn from Federal Reserve Board Statistical Release H.6. The data are available at a monthly frequency, seasonally adjusted, for the period since January 1959. Data on interest rates corresponding to the monetary assets are not as readily available. Our measure of \(r_{\text{OCD},t}\), the interest rate on other checkable deposits, is an average across various types of interest-bearing accounts computed by the Federal Reserve Board. One potential objection to our procedure is that we do not attribute any “implicit” interest on checking accounts. Insofar as depositors are given toasters or other goods, our measure is incorrect. The literature on implicit interest on demand deposits (see Startz 1979), however, has emphasized that banks give transactions services to their customers. The value of these transactions services is captured by the aggregator function \(f\), so no further adjustment is needed in our method to deal with these services.

The interest rate on savings accounts at commercial banks, as well as those on money-market accounts, \(r_{\text{SCB},t}\), \(r_{\text{MMCB},t}\) and \(r_{\text{MMSL},t}\) were also provided by the Federal Reserve Board. We constructed the interest rate on savings accounts at thrifts, \(r_{\text{SSL},t}\), in two parts. Since mid-1986, when interest rates at thrifts were deregulated, we use the average interest rate paid on these accounts as reported by the Federal Reserve Board. Before 1986, we impute interest rates for thrifts using rates at commercial banks. Our imputation relies on regulations relating the maximum allowable interest rates at thrifts and commercial banks, as well as occasional Federal Home Loan Bank Board surveys of average interest rates paid at these institutions. For the 1959–1970 period, we assumed \(r_{\text{SSL},t} = r_{\text{SCB},t} + .75\). This additive factor was .50 for 1970–1973. For the 1973–1979 period we assume \(r_{\text{SSL},t} = 5.24\%\), and for the 1979–1986 period, we make crude estimates using occasional surveys of interest rates paid. For part of our sample, the interest rate on passbook savings accounts exceeded that on commercial paper. We assume that these interest rates were equal in these cases, thereby setting the weight on the associated asset stock to 0.

Figure 1 plots M1 and the CE aggregate based on contemporaneous spreads defined in (15). The two aggregates are drawn on the same scale, so the figure shows that their typical size and low-frequency properties are similar. Although CE includes more assets, many of these are given small weight. In particular, other checkable deposits are given a smaller weight in CE than in M1. The most striking difference between the two aggregates is that M1's rate of growth is smoother. CE rises much more sharply than M1 in some periods, and CE also declines steadily during certain periods. These pronounced changes in CE are associated with changes in interest rates. When the commercial-paper rate changes, the rates on regulated monetary assets often change less than proportionately. Thus, as an example, declines in the commercial-paper rate tend to raise \(r_{\text{OCD},t}/r_{\text{CP},t}\), so that the weight on OCD—and the aggregate itself—declines.

The extreme sensitivity of the CE aggregate to interest rates is somewhat problematic. The reason is that it is hard to believe that the liquidity characteristics of existing assets change by as much as do the ratios \((r_{\text{CP}} - r_t)/r_{\text{CP}}\). Rather, what occurs is that many interest rates on monetary assets are kept relatively constant, whereas the interest rate on commercial paper is more variable. If people readjust their portfolios continually, the CE formula would still be valid but, because agents do not do so, the effect of changes in \((r_{\text{CP}} - r_t)/r_{\text{CP}}\) on actual liquidity is exaggerated.

These high-frequency movements in CE raise the question of whether CE and traditional aggregates exhibit different stationarity properties. This issue is important because part of our subsequent analysis will focus on the long-run effects of monetary shocks on the level of economic activity.

The first panel in Table 1 reports standard augmented Dickey–Fuller tests of the null hypothesis that various mone-
tary aggregates contain unit roots. These tests are equivalent to testing whether $\gamma = 0$ in an equation of the form

$$\Delta \ln M_t = \alpha + \beta t + \gamma \ln M_{t-1} + \sum_{j=1}^{\infty} \delta_j \Delta \ln M_{t-j} + \epsilon_t^M. \quad (16)$$

Following the procedure recommended by Stock (in press), we choose the value of $\kappa$ through the SBIC. This ensures that we do not lose power through underparameterization, by failing to eliminate serial correlation in the residuals, or through overparameterization. These values are reported in the second and fourth columns. The other two columns of the table report tests constraining $\beta$, the coefficient on the linear trend, to equal 0 and also tests allowing it to be unrestricted. As West (1987) suggested, it is more reasonable to allow $\beta$ to be unrestricted.

Although we find that, when $\beta = 0$ the hypothesis that there is a unit root cannot be rejected for any of the aggregates, when $\beta$ is unrestricted this hypothesis can be rejected for currency and several of our CE aggregates (contemporaneous CE, CE-1, and ARIMA CE-3) at the 1% level. This rejection reflects the cyclical patterns in Figure 1 and is probably due to high-frequency interest-rate movements. The rejection might thus disappear in a large sample, which might include more low-frequency variation in CE.

Since even augmented Dickey–Fuller tests have low power, we also present tests of the alternative hypothesis—namely, that the variables are stationary. The last two columns of Table 1 report Lagrange multiplier (LM) statistics from a test derived by Kwiatkowski, Phillips, Schmidt, and Shin (1992) of the null hypothesis $\sigma^2 = 0$ in the specification

$$y_t = \xi t + r_t + \epsilon_t, \quad \epsilon_t \sim \text{iid} N(0, \sigma^2),$$

$$r_t = r_{t-1} + u_t, \quad u_t \sim \text{iid} (0, \sigma^2).$$

The test is performed by regressing the variable of interest on a constant or a constant and time trend, taking the residuals (denoted $e_t$) from that regression, forming the sequence of partial sums $\{S_t\}$ of those residuals, and forming the statistic

$$T^{-2} \sum S_t^2/\sigma^2, \quad (17)$$

where $\sigma^2$ is a consistent estimator of the variance of the $\epsilon$ process. For the latter, we use a nonparametric kernel-density estimator from Andrews (1991) that uses a quadratic spectral kernel. That article established that the latter ensures a positive definite value of $\sigma^2$ and provides a procedure for selecting the kernel bandwidth as a function of the sample size. Kwiatkowski et al. (1992) showed that this statistic converges to a functional of Brownian motion free of nuisance parameters and used Monte Carlo techniques to obtain critical values.

Although we can reject the null hypothesis of level stationarity for all the aggregates, we cannot reject the null of trend stationarity for any of the CE aggregates at the 5% level. This suggests that if there is a difference-stationary component to the CE aggregates it is small; again, these results are probably due to the high-frequency movements in interest rates.

Both sets of results leave us with a choice as to how to treat the stationarity of CE. We wish to compare the properties of CE with those of the traditional aggregates. Conventional specifications—for example, those of Kormendi and Meguire (1984) and Stock and Watson (1989)—treated the latter as nonstationary. Our basic specifications will therefore assume that CE has one unit root as well. We will, however, also discuss how the results change when we assume instead that CE is stationary.

### 3. TESTING METHODS

Developing a new monetary aggregate is of some intrinsic interest for what it reveals about the evolution of the money supply in the United States. The new aggregate may also yield new evidence on the longstanding controversy, recently surveyed by Blanchard (1990), on whether monetary shocks affect real activity. Before presenting evidence on this issue,
it is worth discussing whether it makes sense to judge monetary aggregates by their correlation with subsequent changes in output. In principle, there exist three separate kinds of shocks that affect the supply of L. The first, which we label εt, is a deliberate change in the policy of the central bank. The second, which we label ε′t, is a change in the supply of money that represents a normal reaction to other economic shocks. One exogenous shock that is of particular importance in this respect is a change in money demand—that is, a change in the marginal utility of L for given interest rates and output. An increase in money demand will lead to an increase in L if the central bank pursues a policy of stabilizing interest rates. Similarly, the supply of L from the private financial system might respond to exogenous shocks that affect interest rates. Finally, there is a shock that we label ε′t that represents exogenous changes in the financial system. These shocks change the supply of L for a given policy of the central bank. Changes in α represent shocks in this category. Thus Lt can be written as

$$L_t = \Omega(L)\epsilon_t + \Omega'(L)\epsilon'_t + \Omega''(L)\epsilon''_t, \quad (17)$$

where the Ω(L)'s represent polynomials in the lag operator. Variables such as M1, which are imperfect measures of L, are also affected by these three types of shocks, though the shocks to the financial system that affect M1 are different from those that affect L. A change in α that makes NOW accounts more attractive and induces households to shift from demand deposits and savings accounts into NOW accounts will tend to raise the CE aggregate but might well not affect M2.

Traditional models in which money affects output predict only that the increases in the first two disturbance terms in (17) are correlated with subsequent increases in output. The effect of the last disturbance is ambiguous. If this disturbance is predominantly due to policies that stabilize interest rates in response to changes in money demand, it would tend to be negatively correlated with output.

If one can ignore the last disturbance in (17), proper measures of Lt ought thus to be more closely correlated with output than more arbitrary measures such as M1. Moreover, they should be more closely linked to output than variables that are more strongly influenced by the central bank, such as the monetary base, because these variables do not include the effects of changes in financial institutions (the ε′′s) which affect L.

The third set of disturbances in (17) plagues virtually any study of the correlation of changes in money and other variables. One study that avoided this problem was Romer and Romer's (1989) analysis of how output is affected by changes in the Federal Reserve’s intentions. It remains true, however, that monetary aggregates that capture more accurately the first two disturbances in (17) should be closely correlated with output if exogenous monetary disturbances do indeed cause output movements. Thus our main set of empirical results concerns the correlations of various measures of money with subsequent movements in output (and prices).

We study the correlation between changes in various monetary aggregates and two measures of real activity, industrial production and the unemployment rate for married men. The latter is a particularly useful measure of labor-market tightness because, unlike other unemployment rates, it is relatively uncontaminated by decisions to leave the labor force.

When we use industrial production as our measure of activity y, we estimate equations of the form

$$\Delta \ln y_t = \alpha + \delta t + \beta(L)\Delta \ln m_t + \gamma(L)\Delta \ln m_t + \epsilon'_t. \quad (18)$$

The polynomials β(L) and γ(L) are of 12th order, which seems natural given that we are analyzing monthly data. Industrial production enters the equation in logarithmic first differences because, following Nelson and Plosser (1982), we assume that the logarithm of industrial production has a unit root. When we study the behavior of the unemployment rate, we use its level, not its first difference, in (17) because the unemployment rate is more likely to be stationary. Following the suggestion of Stock and Watson (1989), we have included a deterministic trend in this regression even when we are explaining the behavior of the change in industrial production. Our results are not sensitive to the inclusion of this trend variable. Equation (17) assumes that only the logarithmic first difference of money is correlated with output. This is reasonable if money itself has a unit root because, in this case, the monetary shocks of (16) are functions only of current and lagged differences in money.

We test for the absence of correlation between money and output by testing the null hypothesis that γ(L) = 0, and we interpret rejections as suggesting that changes in money affect economic activity. There are many factors, such as nominal rigidities and issues pertaining to the distribution of newly supplied monetary assets, that could make monetary shocks nonneutral in the short run. It is much more difficult, however, to explain permanent, long-run effects of money on output, particularly if these long-run effects are similar in size to the largest short-run effects. Even in models in which monetary shocks have long-run effects, these are typically much smaller than the short-run effects.

In the case in which both output and the monetary indicator have a unit root, our test for monetary neutrality corresponds to the one proposed by Kormendi and Meguire (1984). In this case, (17) can be rewritten as

$$\Delta \ln y_t = \frac{\alpha + \delta t + \gamma(L)\Delta \ln m_t + \epsilon'_t}{1 - \beta(L)}. \quad (19)$$

A permanent 1% increase in money raises long-run output by \((1 - \beta(1))^{-1}\gamma(1)\%\). Testing long-run neutrality therefore requires testing whether the estimated value of this quantity is significantly different from 0.

One question that arises at this point is why our tests of long-run neutrality make sense given the criticism levelled by McCallum (1984) at the earlier attempt by Lucas (1980). As McCallum showed, Lucas's (1980) method is equivalent to computing the sum of the coefficients in a regression of output on many lags of money, and the same is true of our method. McCallum's (1984) criticism of this method applies only when money is stationary. Then, monetary shocks have no long-run effect on money itself. So, as argued by Fisher and Seater (1993), measuring the response to a once-and-
for-all increase in money is then impossible because such increases never occur in practice.

Suppose, however, that money has a unit root. Then a shock that raises money at \( t \) by 1% raises money at \( t + \infty \) by 2%. The quantity \( Z \) can be estimated from univariate regressions of money growth on lagged values of money growth using standard methods. Then the sum-of-coefficients method works because an event analogous to a once-and-for-all increase in money actually takes place every period.

The quantity \( Z \) is generally not equal to 1, but this does not matter for testing long-run neutrality. From (18) and the univariate regressions for money growth, we can compute two things—\( Z \), the long-run response of money to a unit innovation in money growth, and \( (1 - \beta'(1))^{-1}\gamma(1)Z \), the long-run response of output to a unit innovation in money. Thus, under the null that \( Z \) is nonzero, long-run neutrality prevails only if \( (1 - \beta'(1))^{-1}\gamma(1) = 0 \).

A certain form of long-run neutrality, whether monetary shocks have permanent effects, can be tested even when money is stationary. This is not identical to the question of what would happen if money rose and stayed higher forever, but it is closely related. If output is itself stationary, then it is impossible for monetary shocks to have long-run effects. If one accepts that industrial production has a unit root, then analysis of the long-run effects of monetary shocks becomes interesting. For the reasons stressed by McCallum (1984), this cannot be tested by the sum of the coefficients method we used in the case in which both money and output have a unit root. Instead, what is required is a simultaneous analysis of (16) and the equation that is analogous to (18) and relates output growth to the current and past levels of money. Using both of these equations simultaneously, one can compute the impulse response function (IRF) of output with respect to monetary shocks and analyze the long-run response to these shocks.

We carry out these tests for the CE aggregate and for two other variables that have recently been proposed as measures of monetary policy, the federal-funds rate and the spread between the yield on commercial paper and the yield on T-bills. All three of these variables may be stationary. If monetary policy is indeed the predominant force acting on these variables, then their changes ought not to have any long-run effects on output.

The existence of the last set of shocks in (17) has led several researchers—for example, Bernanke and Blinder (1992)—to postulate that money-supply shocks should really be associated with changes in interest rates. If the central bank simply sets interest rates exogenously, then the exact nature of \( L \) is not very important for understanding the effect of monetary policy. The view that interest rates are set exclusively by exogeneous changes in policy is not very tenable on a priori grounds because interest rates have been very volatile during certain periods. On the other hand, this view has gained prominence since Sims’s (1980) demonstration that changes in M1 lose their ability to predict changes in output once one controls for interest rates.

A more plausible view of interest rates is that they are affected by the three shocks in (17) and, in addition, by other shocks such as changes in fiscal policy and the expected future profitability of current investment. The presence of such shocks implies that movements in a proper monetary aggregate ought still to be correlated with subsequent movements in output after controlling for interest rates. It is thus of interest to see whether CE, which supposedly is better at capturing movements in \( \epsilon_t \), maintains its ability to predict output after one controls for interest rates. This test is particularly valuable because changes in interest rates affect the CE aggregate directly.

We also test whether money has any effect on prices by considering regressions of the form

\[
\Delta \ln p_t = \alpha' + \delta_t + \beta'(L)\Delta \ln m_t + \gamma'(L)\Delta \ln m_t + \epsilon_t, \tag{20}
\]

where we measure the price level \( (p_t) \) as the Consumer Price Index (CPI). We then test the null hypothesis that money has no effect on prices by testing whether \( \gamma'(L) = 0 \). Admittedly, there are no plausible theories that imply this null hypothesis. It is thus of more interest to analyze the long-run response of prices to monetary shocks. As shown by Fisher and Seater (1993), this can be done using the same method that we used for measuring the long-run effect on output. In the case in which money and prices both have a unit root, the hypothesis that permanent increases in money eventually raise prices one for one implies that \( (1 - \beta'(1))^{-1}\gamma'(1) = 1 \).

Finally, we consider VAR’s that are analogous to those of Sims (1980). VAR’s have been proposed as a solution to the problem posed by the last set of shocks in (17). The idea, developed by Sims (1986), Bernanke (1986), and Blanchard (1989), is to orthogonalize the innovations affecting the series included in the VAR in such a way that one isolates a component that is associated mainly with \( \epsilon_t \), another that is related to shocks to financial markets, and another that is related to policy shocks. Many different decompositions of this type have been proposed.

VAR’s provide additional evidence on the role of monetary aggregates once one controls for other variables. They are also of interest because they allow us to check the robustness of our conclusions regarding short- and long-run neutrality to the inclusion of other variables. Because the other included variables are unlikely to be exogenous, it is not clear that these tests of neutrality have a structural interpretation.

4. MONEY AND OUTPUT

4.1 Granger-Causality Tests

Before examining whether any of the monetary aggregates or indicators Granger-causes (or, more accurately, provides incremental explanatory power in forecasting) various measures of real activity, we must examine the long-run relationships between those variables. If, in fact, a monetary aggregate is cointegrated with a measure of real activity, then there must be a Granger-causal ordering between those two variables. Cointegration would also imply that our bivariate
systems are misspecified: They should, in fact, be respecified as error-correction models.

Since unemployment and the interest-rate variables are stationary, such issues are only relevant to the case of industrial production and the various monetary aggregates. We test for cointegration between industrial production and each of the aggregates using the standard technique described by Engle and Granger (1987), in which the unit-root test for the residuals of the first-stage regression is done using an augmented Dickey–Fuller test with lag length selected by the SBIC, and find that none of the pairs of variables are cointegrated.

We first report tests of whether money matters at all in equations such as (18). The first eight rows of Table 2 report the rejection $p$ values for the null hypothesis of $\gamma = 0$—that is, of no link between current real activity and past money for a range of traditional monetary aggregates. The results do not yield a consistent pattern except for M2, which always has a statistically significant effect on real activity. When the measure of activity is industrial production, the null of monetary neutrality is rejected at high levels for both ordinary and Divisia M2 and the Divisia M3 aggregate proposed by Barnett et al. (1992) and Barnett and Yue (1991) but not for M1 or M3. The ease with which we can accept that M1 has no effect on industrial production contrasts with Stock and Watson (1989), who showed that M1 does affect output in specifications like (18). The reason our results are different is that our sample is longer than theirs. We obtain their result when we restrict attention to the period before 1985.

In contrast to the results for industrial production, in the case of the unemployment rate, we obtain statistically significant rejections for M1 and M2, but neither Divisia M2 nor M3, nor ordinary M3, yield strong rejections of neutrality. Other conventional monetary aggregates yield even weaker evidence against neutrality. For currency and the money multiplier, we do not reject the null of monetary neutrality for either industrial production or the unemployment rate. For the monetary base, the null is rejected when we focus on the unemployment rate, but not when we consider industrial production.

We report next the results for six versions of the CE aggregate. The first is the one using current interest-rate spreads as weights; we label this the contemporaneous CE. Next, we look at CE-1 and CE-3, which use one- and three-year moving averages of actual spreads. Then, we study ARIMA CE-1 and ARIMA CE-3, which use forecasts of future spreads. Finally, we study the fixed-weight CE, which uses the overall sample average yield spreads as weights.

The CE aggregates consistently help predict industrial production and, in the majority of cases, help predict unemployment as well. Thus the way in which we construct the weights does not seem all that important in obtaining aggregates that are more strongly correlated with subsequent changes in economic activity than are the more traditional aggregates. In what follows, we will concentrate mainly on the effects of ARIMA CE-3 for two reasons. First, it is relatively smooth—that is, insensitive to short-term changes in interest rates. Second, aside from the coefficients in the ARIMA process for interest-rate spreads, it uses only lagged information.

The remaining rows of Table 2 present Granger-causality tests for several recently proposed alternative measures of monetary policy. Bernanke (1990) argued that shifts in policy are captured by the yield spread between commercial paper and T-bills. The correlation of this spread with subsequent output was first demonstrated by Friedman and Kuttner (1990). We too find strong correlations with output. We find similarly strong rejections of $\gamma(L) = 0$ in (18) when we use the federal-funds rate suggested by Bernanke and Blinder (1990) or the commercial-paper rate itself as a measure of monetary policy.

One important limitation of interest-rate-based measures of monetary stance is that the estimated effects of these aggregates are sensitive to the time period in question. Using the CE aggregate, within-sample stability tests for the coefficients in (18) or industrial production, for example, do not reject the null hypothesis that $\gamma(L)$ and $\beta(L)$ are identical for the pre- and post-1980 period. By contrast, subsample stability is rejected at the .01 level for the commercial-paper–T-bill yield spread. The evidence against subsample stability for M1 and M2 is also weak, suggesting that these aggregates behave more like the CE aggregate.

The results in Table 2 suggest that the CE aggregate provides stronger and more consistent evidence against the mon-

<table>
<thead>
<tr>
<th>Table 2. Causality Tests: Monetary Aggregates and Real Activity</th>
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<tbody>
<tr>
<td><strong>Traditional monetary aggregates</strong></td>
</tr>
<tr>
<td>M1</td>
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<tr>
<td>M2</td>
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<tr>
<td>M3</td>
</tr>
<tr>
<td>Monetary base</td>
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<tr>
<td>Money multiplier</td>
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<tr>
<td>Currency</td>
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<tr>
<td>Divisia aggregates</td>
</tr>
<tr>
<td>Divisia M2</td>
</tr>
<tr>
<td>Divisia M3</td>
</tr>
<tr>
<td>Alternative indicators of monetary policy</td>
</tr>
<tr>
<td>Commercial paper</td>
</tr>
<tr>
<td>Federal-funds rate</td>
</tr>
<tr>
<td>Commercial-paper–T-bill yield spread</td>
</tr>
<tr>
<td>Currency equivalent aggregates</td>
</tr>
<tr>
<td>Contemporaneous CE</td>
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<tr>
<td>CE-1</td>
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<tr>
<td>CE-3</td>
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<tr>
<td>ARIMA CE-1</td>
</tr>
<tr>
<td>ARIMA CE-3</td>
</tr>
<tr>
<td>Fixed-weight CE</td>
</tr>
</tbody>
</table>

NOTE: Each entry reports the confidence level at which the null hypothesis that $\gamma(L) = 0$ can be rejected in the specification $\Delta \ln y_t = \alpha + \beta t + \beta(L) \Delta \ln y_t + \gamma(L) \Delta \ln m_t + \epsilon_t$. Both $\beta(L)$ and $\gamma(L)$ are 12th-order lag polynomials, so the statistics reported in the table test the joint hypothesis that $\gamma_1 = \gamma_2 = \cdots = \gamma_{12} = 0$. The sample period for all estimates except Divisia M2 and M3 is 1960:2–1989:6. For the two Divisia aggregates, the sample is 1961:2–1989:6.
etary neutrality hypothesis than many, but not all, other measures of the money supply. The interest-rate variables, however, are actually better predictors of output. Since CE is partially based on interest rates, this raises the question of whether CE has any incremental explanatory power once the other measures of money or monetary policy are included in an equation for output.

Table 3 addressed this question by reporting p values for the F tests of the hypothesis that all coefficients on the lagged values of ARIMA CE-3 are 0 in equations that already include other indicators of monetary stance. The findings suggest that CE’s information content remains even after the other measures of money or monetary policy are controlled for. The most important finding in the table is that CE remains a significant explanator of industrial production, though not of the unemployment rate, after controlling for interest rates. This suggests that the variations in interest rates that affect the benchmark CE are not solely responsible for its explanatory power.

This finding also provides a sharp contrast with results for simple sum aggregates. The last two rows of Table 3 provide test statistics for the additional explanatory power of M1 and M2 once one controls for the commercial-paper rate. As suggested by Sims (1980), interest rates rob traditional monetary aggregates of much of their explanatory power. The relationship between M1 and our two activity variables ceases to be statistically significant at the 5% level; M2 remains significantly correlated only with unemployment.

### 4.2 Long-Run Neutrality

We now test whether once-and-for-all increases in money change output significantly in the long run. We thus measure 

\[ (1 - \beta(1))^{-1} \gamma(1) \]

the percentage increase in long-run output when money increases by 1%, and test whether it equals 0. Table 4 displays this measure and its standard error for conventional, CE, and Divisia aggregates.

The results provide strong evidence against long-run neutrality for M1, M2, and both the Divisia M2 and M3. Not only is the hypothesis that 

\[ (1 - \beta(1))^{-1} \gamma(1) = 0 \]

evaluated at \( L = 1 \), where the lag polynomials are estimated as in Table 2. Standard errors, which are shown in parentheses, are calculated by Poterba, Rotemberg, and Summers (1986) as 

\[ \hat{\sigma}^2 = \hat{\sigma}^2 + \hat{\sigma}^2 \hat{\sigma}^2 \]

where \( \hat{\sigma}^2 \) is the variance–covariance matrix of the estimated parameters.

The null cannot be rejected for the CE aggregate nor for its three-year and contemporaneous-weight variants. For the benchmark CE, the point estimate of the long-run elasticity of output with respect to money is very small (0.05). For the contemporaneous and three-year versions, the point estimates are somewhat larger but still below 0.15. On the other hand, the long-run response of output to a 1% change in the fixed-weight CE is similar to that of output to M1. The point estimate of this response is near 1, and it is significantly different from 0. The utility-based aggregates that do not let asset characteristics vary over time, such as the fixed-weight CE and the Divisia aggregates, do not appear to be long-run neutral. This suggests that changes in the liquidity characteristics of assets may be important.
4.3 Impulse-Response Functions

In this subsection we present IRF's of output with respect to monetary shocks. We do this first for the monetary indexes that are arguably nonstationary. For these we report the effects of a once-and-for-all increase in money using the estimates of (18). From these we learn the time profile of the effects of these monetary impulses. We then study the IRF's for the indicators of monetary policy that appear stationary. These IRF's, which are based on simulating the analogs to Equations (16) and (18) simultaneously, also serve to test the long-run neutrality of these shocks.

Figure 2, a–c, displays these IRF's for the effects on industrial production of a 1% once-and-for-all increase in M1, M2, and CE. The figure also reports one standard-error band on either side of the point estimates. Each part of the figure shows the cumulative effect of a shock to the monetary variable on real output. Long-run neutrality therefore implies a zero effect at long horizons.
The IRF's of industrial production with respect to M1 and M2 do not correspond to the predictions of standard models. The cumulative effect of money on output rises for the first 14 months after a monetary shock. There is virtually no reduction in the cumulative output effect thereafter, which leads to the finding in Table 4 of strong evidence against the long-run neutrality hypothesis.

The IRF pattern is substantially different for the CE aggregate. A positive innovation in the money stock raises industrial production, and the cumulative IRF continues to increase for 7 months after the shock. After this point, however, the lagged monetary shock has a negative effect on output, and by 11 months after the shock the total increase in real output is 0. After that, output continues to decline, but this effect is not statistically significant.

We now turn to the effect of the stationary monetary impulses. The IRF's for shocks to the commercial-paper–T-bill spread and to the federal-funds rate are given in Figure 2, d and e. Figure 2d shows that a 100-basis-point increase in the spread between the annualized yield on commercial paper and that on T-bills permanently lowers industrial production by 5%. It is hard to interpret this as a response to a monetary shock because the maximal reduction in output, which occurs after about 10 months, is essentially equal to the long-run reduction. Figure 2e shows a similar pattern for the response of output to a 100-basis-point increase in the annualized federal-funds rate. Output falls steadily and becomes about 5% lower in the long run. This too seems hard to reconcile with typical analyses of monetary contractions.

We also computed the impulse response of industrial production to shocks in CE assuming that CE is stationary. In other words, we considered the system of Equations (16) and (18), in which CE enters only in levels, and computed the response of both variables to $\epsilon_n$. We did this because the results of Table 1 suggest that the CE aggregate might be stationary. The results are very similar to those obtained when CE is assumed to have a unit root. Increases in CE first raise output significantly and then lower output so that they have no significant long-run effect.

Because increases in the commercial-paper rate tend to lower the CE aggregate, these results can be seen as puzzling. The simple correlation between the level of the contemporaneous CE aggregate and the commercial-paper rate is .77. How can increases in contemporaneous CE, which are largely associated with increases in interest rates, lead to high output when the opposite is true of interest-rate increases on their own? The answer is partly that the variables are not identical and partly that interest-rate increases do, initially, raise output. As can be seen in Figure 2e, this initial increase is both small and statistically insignificant. The incorporation of interest variations into the CE aggregate might thus have the effect of magnifying this initial positive response.

5. MONEY AND PRICES

The CE aggregate can also be used to examine the effects of monetary shocks on the price level. In the long run, if output is not affected by monetary shocks and velocity is stable, prices should rise one-for-one with increases in the money stock. This section tests whether monetary impulses Granger-cause changes in the price level and investigates the long-run price-level effects of monetary shocks.

The second column in Table 5 reports tests of Granger-cause between monetary aggregates, other indicators of monetary policy, and the price level. The entries correspond to p values for rejecting the null hypothesis that $\gamma(L) = 0$ in Equation (20). All of the CE aggregates except that with time-invariant weights display strong predictive power for future movements in the price level. The Granger-cause results are much weaker for other aggregates. The monetary base is the only such aggregate for which we can reject the null hypothesis of no effect. The last three rows of the table show that the federal-funds rate and the commercial-paper rate do show strong predictive power for changes in the price level, but the null hypothesis of no causality cannot be rejected for the commercial-paper–T-bill yield spread.

The third column in Table 5 presents our estimates of the long-run effects of monetary shocks on prices. The results

<table>
<thead>
<tr>
<th>Traditional aggregates</th>
<th>Causality</th>
<th>Long-run effect</th>
</tr>
</thead>
<tbody>
<tr>
<td>M1</td>
<td>.241</td>
<td>2.42</td>
</tr>
<tr>
<td>M2</td>
<td>.394</td>
<td>.86</td>
</tr>
<tr>
<td>M3</td>
<td>.693</td>
<td>.95</td>
</tr>
<tr>
<td>Monetary base</td>
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<tr>
<td>Money multiplier</td>
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<td>Currency</td>
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<td>1.52</td>
</tr>
<tr>
<td>Divisia aggregates</td>
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<td></td>
</tr>
<tr>
<td>Divisia M2</td>
<td>.802</td>
<td>.94</td>
</tr>
<tr>
<td>Divisia M3</td>
<td>.640</td>
<td>1.21</td>
</tr>
<tr>
<td>Currency-equivalent aggregates</td>
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<td></td>
</tr>
<tr>
<td>Contemporaneous CE</td>
<td>.046</td>
<td>.31</td>
</tr>
<tr>
<td>CE-1</td>
<td>.002</td>
<td>.62</td>
</tr>
<tr>
<td>CE-3</td>
<td>.002</td>
<td>.81</td>
</tr>
<tr>
<td>ARIMA CE-1</td>
<td>.029</td>
<td>.35</td>
</tr>
<tr>
<td>ARIMA CE-3</td>
<td>.002</td>
<td>.71</td>
</tr>
<tr>
<td>Fixed-weight CE</td>
<td>.141</td>
<td>3.71</td>
</tr>
</tbody>
</table>

NOTE: Entries in column 2 are p values for Granger-cause tests of the sort reported in Table 2, but entries in column 3 are long-run responses of the price level to shocks in the growth rate of the monetary aggregates, whose analogue for industrial production was reported in Table 3. Standard errors are in parentheses.
do not permit us to draw sharp conclusions about the relative merits of the various monetary aggregates. For all of the monetary aggregates we consider, the hypothesis that in the long run prices rise in proportion to the rise in money cannot be rejected at conventional significance levels. The standard errors are large, however, and the point estimates for the long-run elasticity of prices with respect to money range from .3 to 3.7. Figure 3, a and b, reports estimates of the CPI’s response to a permanent 1% increase in M1 and CE. In the case of CE, half of the long-run effect (.62) is reached three years after the monetary shock. This implies that the process by which prices respond to monetary shocks is relatively slow and gradual.

6. VECTOR AUTOREGRESSIONS

The last two sections have presented bivariate evidence on the relationship between money, output, and prices. We now consider the robustness of our finding that the CE aggregate is the best predictor of changes in output. In particular, we check whether this finding holds up in VAR’s that include the four variables of Sims (1980). Listed in the order in which they appear in the VAR, the variables are the commercial-paper rate, a monetary aggregate, the CPI, industrial production, and unemployment. Except for the interest rate, which is in levels, the variables are in logarithmic first differences. The ordering is the same as that of Sims (1980) except that we have added the unemployment rate at the end.

Figure 4, a and b, gives the impulse responses of all the variables to all five shocks for the cases of M1 and ARIMA CE-3. These impulse responses are constructed cumulatively so that they give the total effect on each variable of each shock. Both monetary impulses permanently raise prices and temporarily raise output and lower unemployment. There are differences, however, in the critical value at which one can reject that each particular monetary aggregate has no effect on industrial production and unemployment. These critical values are .642 for M1, thus confirming the Sims (1980) finding, and .014 for ARIMA CE-3. The critical value for the analogous VAR with M2 is .032.

The difference between the two monetary aggregates is less pronounced in these results than in the bivariate regressions. The effects of both CE and M1 on output are now temporary. Moreover, the two IRF’s have in common that an increase in the monetary aggregate raises the interest rate, though this effect is smaller for ARIMA CE-3. This positive effect of monetary shocks on interest rates suggests that these innovations are not due exclusively to changes in \( \epsilon_t^m \) and \( \epsilon_t^i \). The reason is that these shocks would normally depress interest rates as they increase the money supply. It is very likely, however, that VAR’s face difficulties in estimating the response of interest rates to monetary shocks. The reason is that interest rates probably reflect anticipations of future changes in the stock of money. For this reason, we find the bivariate response reported earlier more reliable.

7. CONCLUSION

This article develops a new monetary aggregate, the CE aggregate, based on the amount households spend on liquidity services. This aggregate resembles Divisia monetary aggregates in that different monetary assets are added together with different weights. Deriving the CE aggregate requires a strong assumption of separability in the liquidity service function, but this assumption yields more robustness with respect to changes in asset characteristics and also permits simple-sum aggregation across households. We also compare the performance of the CE monetary aggregate, other traditional monetary aggregates, and various indicators of monetary policy in forecasting real activity. The CE aggregate has more predictive power than any of the traditional monetary aggregates. Although some other indicators of monetary policy based on nominal interest rates can forecast output as well as the CE aggregate, they also suggest long-run nonneutrality of output with respect to money. Changes in the CE aggregate are therefore a more plausible indicator of changes in the supply of liquid assets.

Our derivation of the CE aggregate yields a simple relationship between this aggregate, consumption, and interest rates. This could in principle provide a basis for studying liquidity demand. Because our analysis does not explicitly allow for costs of adjustment when households and firms change their holdings of liquid assets, however, there are reasons to suspect that empirical work based on this liquidity-demand equation would fail. In our exploratory analysis of this issue, we estimated a log-linear liquidity-demand equation relating
Figure 4. Impulse-Response Function a: With M1 and b: With ARIMA CE-3.
the logarithm of CE to a time trend, the logarithm of real GNP, and the commercial-paper rate. The estimated coefficient on the commercial-paper rate was positive, contrary to the theory's prediction of a negative effect. Further work could explore the demand for liquidity services in models that provide a more realistic description of dynamics.

Our results also contain some information on the channels through which monetary disturbances affect output. The literature on this topic stresses two issues, whether nominal rigidities are important and whether money or credit is important. The distinction between the money and credit views has been stressed mainly by authors working with models of sticky prices, including Bernanke and Blinder (1990), Gertler, Hubbard, and Kashyap (1990), Kashyap, Stein, and Wilcox (1993), and Romer and Romer (1990). This distinction is, however, just as important if nominal rigidities are absent and money matters only because different people carry out financial transactions at different times. This can be seen by comparing the work of Grossman and Weiss (1983), Rotemberg (1984), and Fuerst (1990).

According to the money view, monetary contractions initially reduce money balances and interest rates must rise to make people hold the new constellation of assets. This increase in interest rates then lowers investment. According to the credit view, there is a direct reduction in investment that comes about because monetary contractions force banks to curtail the volume of loans. Our findings provide some support for the money view because they suggest that the money stock that has the most plausible effect on output is one that makes the most sense from the point of view of holders of money. A monetary expansion can thus be thought of as a situation in which the level of liquid assets is high, at least for some consumers.

In the credit view, the level of consumers' liquid assets does not matter, except insofar as it is correlated with funds available for loans. The latter are presumably more correlated with broad aggregates like M2 than with the CE aggregate. From the point of view of credit availability, whether banks raise their funds through "regular" checking accounts at zero interest or through expensive certificates of deposit is probably irrelevant. Because the data suggest that this distinction is important for the subsequent evolution of output, they also suggest that the credit view is incomplete.

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